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Abstract

Building on anthropological evidence, we develop a model of intra-household decision making on fertility and child survival within the framework of the collective household model. We carry out a test of the implications of this framework with data from Demographic and Health Surveys in rural Mali, where polygyny rates among married women are close to 50 per cent. The econometric tests reject the implications of efficient intra-household allocations for junior wives in bigynous households and fail to reject for senior wives in bigynous households as well as for wives in monogamous households. These findings are consistent with existing narrative evidence according to which co-wife rivalry is responsible for resource-consuming struggle and junior wives are the adults with the weakest bargaining position in the household.

1 Introduction

Over the past twenty years, both the theoretical and empirical literature on intra-household decision making has grown rapidly. A major concern of this research has been whether resources within a household are allocated efficiently. On the theoretical side, the unitary model (Becker, 1994), Nash bargaining (Manser and Brown, 1980; McElroy and Horney, 1981; Lundberg and Pollak, 1993), and, more recently, the collective model (Chiappori, 1988; Bourguignon and Chiappori, 1992) are well-studied approaches that predict Pareto-efficient intra-household allocations. On the other hand, certain non-cooperative bargaining models (Kanbur and Haddad, 1994) generate outcomes that are not in general Pareto-efficient. On the empirical side, Bourguignon et al. (1993), Browning et al. (1994), and Browning and Chiappori (1997) test implications of the collective model and, using data from Canada, fail to reject the hypothesis that intra-household allocations are Pareto efficient. The unitary model, on the other hand, is rejected throughout. Similarly, with data from various developing countries, Thomas and Chen (1994) and Mallucio and Quisumbing (2003) also reject the unitary model, but not Pareto efficiency. On the other hand Udry (1996), who compares farm yields of field plots cultivated by different members of the same household in Burkina Faso, finds significant differences in yields depending on the gender of the cultivating household member and thus rejects intra-household efficiency. Similarly, Duflo and Udry (2005), who investigate changes in household expenditures in response to rainfall shocks which affect agricultural output on husbands' and wives' field plots differently, reject the collective model with data from Cote d'Ivoire.

This paper makes two novel contributions to this literature. First, we develop a theoretical framework of investment into children by parents, which delivers a test for efficient intra-household allocations. Second, we conduct the efficiency test for two distinct sets of households, monogamous and polygynous ones, which delivers new insights into the intra-household economics of polygyny.

In the theoretical part of the paper, we build on anthropological evidence from rural Mali to develop a model of efficient investments in sons and daughters by parents. We characterize the pattern according to which survival chances of a wife's sons and daughters co-move in response to changes in different household characteristics when intra-household allocations are efficient. In the empirical part, we test for violations of this pattern with Demographic and Health Survey data from rural Mali. Our efficiency test rejects for junior wives in bigynous households while it fails to reject for senior wives as well as for wives in monogamous households. These findings are in accordance with existing qualitative evidence according to which the offspring of a junior wife, who is the adult with the weakest position inside a household, is most severely affected by co-wife rivalry and competition for own surviving offspring (Strassmann, 1997).

Our approach to testing efficiency is related to Udry's (1996) in spirit. He compares input intensities between men's and women's farm plots within the same household and finds that women's plots are cultivated less intensively, which contradicts efficient intra-household factor allocations. Unlike Udry's, however, our test does not require a comparison of allocations across adult members of the household. Instead we focus on differences in mortality outcomes between sons and daughters of a given wife and exploit the fact that the sex of a child is randomly assigned to her. Our test for efficient allocations in polygynous households thus also differs from the one proposed by Mammen (2004). She focuses on differences in child outcomes across co-wives and does not reject a variation of the collective model with data from Cote d'Ivoire.

The remainder of this paper is structured as follows. The next section introduces the data, discusses theories of polygyny, and summarizes relevant anthropological evidence. Section 3 develops a theory of fertility and efficient investment into children, and derives testable implications. The econometric specifications are derived in section 4. The empirical results are presented in section 5. The final section summarizes and concludes.

2 Polygyny in Rural Mali

In this section, we introduce the data used in the empirical analysis and provide some background on polygyny. The exposition of the latter is divided into a discussion of general theories of polygyny and a review of micro-level anthropological evidence.

Demographic and Health Survey Data

The empirical analysis is based on data from the 1996 and 2001 waves of the Demographic and Health Survey (DHS) in Mali, one of the world's poorest and least developed countries. By way of background, Mali ranks 174 out of 177 in the 2003 Human Development Report[1]. We choose this country for three reasons. First, there is a comparatively high incidence of polygyny. With 42.6% of married women having at least one co-wife in 2001, Mali ranks sixth in Tertilt's (2005) country polygyny list[2]. Second, child mortality (death before completing five years of age) in Mali ranks fifth in the world and is second only to Niger among countries that do not currently

[1]Only Burkina Faso, Sierra Leone, and Niger have a lower score on the human development index.

[2]The incidence of polygyny is higher in Cameroon, Burkina Faso, Guinea, Senegal, and Benin.

experience a war or civil war. Finally, there are two relatively recent waves of DHS data available with comparatively large sample sizes, 9704 and 12817 women, respectively.

We further restrict our sample to rural households for two reasons. First, the vast majority of respondents resides in rural areas (72.8 per cent in 2001). Second, the rural data exhibit considerably more variation in the variables of interest for this study, polygyny and child mortality. In 2001, 45.5 per cent of married women lived in polygynous unions in rural areas compared to 33.6 in urban areas. The child mortality percentages are 25.3 for rural and 18.4 for urban areas.

The DHS is nationally representative and the sample selection uses a two-stage stratified random sampling design. The survey collects information on household and individual characteristics, including access to health and education facilities, literacy and education levels of household members, as well as other individual characteristics. Individual information is collected for women aged between 15 and 49 years and for men aged between 15 and 59 years. For women, information on fertility history and antenatal care for each pregnancy is recorded. Although information on income is not collected, data on household possession, such as bikes and motor bikes, electronic devices such as radio receptors, the type of the roof, the walls and the floors of the buildings a household lives in are recorded.

To eliminate potential heterogeneity due to varying intensities of polygyny, we restrict the core sample of our analysis to households with exactly one or two wives at the time of the interview. By way of background, 71 per cent of sampled polygynous households are bigynous while the remaining 29 per cent comprise three or more wives.

Table 1 summarizes the data. Fertility is high, as reflected by 5.6 children ever born to each woman (unadjusted for age). The average age at the time of interview is 32 years and the time spent in marriage about 16, which implies that females marry at an age of 16 on average. Between the two waves the incidence of polygyny has slightly declined. In 1996, 46% of rural women had at least one co-wife compared to 45% in 2001.

Given the censored nature of the data - at age 32, the average age at the interview, most women are still reproductive - we use spacing between births as an inverse measure of fertility. According to Table 1, monogamous wives have an average birth interval of 30.5 and bigynous wives of 31.3 months, implying that women in monogamous households are more reproductive.

As measure of child health outcomes we focus on child mortality, which is defined as death during the first five years of age[3]. At 26.7 per cent child mortality is very high by international

[3]While data on some health care inputs like vaccination and medical treatment by a nurse or doctor are available and would in principle be of interest for this study, all of those measures exhibit very little variation and are thus not suited for the subsequent econometric analysis.

standards[4]. Unlike in some south Asian countries, there is no immediate evidence for a son preference, as mortality is 2.3 percentage points higher for boys than for girls. It is, however, remarkable that this difference is only 1.7 in monogamous households, -0.2 for children of senior bigynous wives and 3.3 for junior bigynous wives.

We now briefly discuss some other variables of interest. In bigynous households, senior wives are almost seven years older than, and junior wives of about the same age as, monogamous wives. Given that both groups are of roughly the same average age at the time of marriage, it follows that second wives are taken approximately seven years later than first wives. At 9.3 percent, female literacy is the same among junior bigynous and monogamous wives while it is smaller by 2.8 percentage points for senior bigynous wives. At least part of this difference is explained by the higher age of senior wives, as female literacy increased on average by a little less than 2 percentage points between the two DHS waves. Male literacy rates are about twice that of females[5]. In our data, bigynous husbands are uniformly less literate than monogamous ones with a difference of 3 percentage points, which, again, is partly explained by the higher age of bigynous husbands. On the other hand, bigynous households hold more wealth on average, as would be expected if each wife contributes to the stock of household assets or brides are ”normal goods” for men.

THEORIES OF POLYGYNY

Since the overall ratio of women to men reaching adulthood in Mali is roughly equal to unity, one might ask how polygyny can be sustained. We will discuss three theories that have been developed in this connection. First, as argued by Becker (1994), heterogeneity in wealth held by males can explain polygyny. If wives are normal goods, wealth inequality among men will lead to polygyny in equilibrium. Provided that women (or their families) have reservation utility in marriage, men who cannot provide the reservation utility are outbid for wives by their wealthier peers. As a result some men take more than one wife while other men remain unmarried. In line with this argument is the economic contribution of women in Sub-Saharan Africa (Boserup, 1970). To the extent that female labor complements physical capital such as land, the shadow price of a wife is relatively smaller for men with a large endowment of physical capital, making them demand more wives on average, which is exactly what Jacoby (1995) finds in an empirical study with data from Cote d’Ivoire.

[4] Notice that this figure is uncorrected for right-censoring. Counting only children that were born at least five years before the interview, would result in an even larger figure.

[5] The DHS does not contain interview data on husbands of all married women. Therefore the number of observations in the husband sample is smaller than if all adult members of all households had been interviewed.

Second, the marriage market squeeze argument has it that there are fewer men than women in the marriage market and that polygyny helps equalizing supply and demand. Three reasons for such a squeeze have been identified (Dorjahn, 1959). First, in a growing population, social arrangements that make men marry later in life than women are responsible for an excess supply of brides relative to grooms. Second, lower life expectancy and higher age at marriage of men causes frequent early widowhood, which increases the supply of brides. Third, a female-biased sex ratio within cohorts is commonly due to higher child mortality among boys and, in rural areas, frequent outmigration by young male adults. In rural Mali, in fact, the population has grown by 1.9% per year between 1995 and 2000 (source: World Bank, 2004) and, according to the DHS data, husbands are on average 13 years older than wives.

Finally, while the marriage squeeze argument takes population dynamics as exogenous, Tertilt (2005) develops a dynamic model in which fertility as well as transfers at marriage are endogenous. Under the assumptions that men make fertility decisions and younger brides are more desirable, she characterizes an equilibrium with a high incidence of polygyny, high fertility, and substantial bride prices (a transfer from the groom to the bride's parents at the time of marriage). Moreover, economic growth is substantially slower than in a regime where polygyny is banned. The reason for high fertility and slow economic growth is that high equilibrium bride prices make daughters a more attractive investment for household heads than other forms of investment, most notably physical capital. A monogamous regime, on the other hand, drives down bride prices, reduces the return of daughters relative to other forms of investment, and thus lowers fertility while fostering faster economic growth.

The approach taken in the present paper differs from the previous ones as we take the structure of households, monogamous versus polygynous, as given and test for efficiency of intra-household allocations. While we are first to conduct a micro-level analysis of polygynous households from an economic angle, there is a body of anthropological literature on the structure and intra-household dynamics of polygynous unions, which will be reviewed next.

Anthropological Evidence

In this subsection, we review a body of literature on the nature of intra-household relations in countries of Sub-saharan Africa. Our focus is on decision making about fertility, parental child rearing inputs, co-wife cooperation and conflict, and the impact of polygyny on wife and child well-being. This will serve as a basis for the theoretical model of fertility and investment in children in the next section.

Of most interest for our study is Strassmann's (1997) account of polygyny and child mortality among Mali's Dogon, a tribe of agripastoralists. The Dogon have patrilineal descent and inheritance, and patrilocal residence. Related women do not marry into the same patrilineage,[6] which makes alliances among female kin difficult and curtails a wife's bargaining position within marriage. Fertility is controlled by the husband. Clitoridectomization is universal, which is meant to reduce female sexual pleasure and promote paternity certainty, and women have to attend menstrual huts, which allows nearly perfect monitoring of the female fertility cycle. According to behavioral scan data, husbands in both monogamous and polygynous unions do very little direct child care. Instead, Strassmann conjectures, males prefer to devote resources to mating effort and consumption. First wives generally have a higher social status and are granted certain non-material privileges, such as having a sleeping room next to the husband's, but no significant material advantage. Each co-wife typically operates her own kitchen and, in some cases, resides in a separate compound. The inheritance and land partitioning system frequently induces co-wife rivalry, which adversely affects child survival, especially boys'. Strassmann elaborates:

> "Cowives are not related, and the rivalry among them extends to their sons, who, upon the death of their father, almost invariably stop farming together. In addition to accusations of neglect and mistreatment, it was widely assumed that cowives often fatally poisoned each other's children. I witnessed special masked dance rituals intended by husbands to deter this behavior. Cowife aggression is extensively documented in Malian court cases with confessions and convictions for poisoning. These cases raise the possibility that Dogon sorcery might have a measurable demographic impact - a view that is consistent with the extraordinarily high mortality of males compared with females. Males are said to be the preferred targets because daughters marry out of the patrilineage whereas sons remain to compete for land. Even if women do not poison each other's children, widespread belief in the hostility of the mother's cowife must be a source of stress. Stressful family environments [...] have been shown to affect childhood cortisol levels and can lead to immunosuppression and a high frequency of illness (Flinn and England, 1995)."

Strassman goes on to discard the view that higher child mortality in polygynous unions is due to resource dilution. Instead she finds that per capita wealth in monogamous and polygynous households does not differ significantly. She also discards the hypothesis of increased health hazards due to crowding in the household (Isaac and Feinberg, 1982).

[6] I.e. there is no sororal polygyny.

Most of Strassmann's observations which are important for the present analysis are confirmed by other recent literature on polygyny and intra-household processes more generally. In a cross-continental comparison, Desai (1992) reports that fathers' involvement in child rearing is generally low in Sub-saharan Africa, especially for children's primary needs. Instead, fathers show more involvement for less vital needs of their offspring, such as education.

The result that polygyny is a risk factor for child survival and health, even when other factors are controlled for, also obtains in two anthropological studies in Tanzania (Hadley, 2005; Sellen, 1999), and in a more aggregate econometric analysis of DHS data from six west African countries (Arney, 2002).

Another, more qualitative, study that documents co-wife conflict among the Dogon is Calame-Griaule (1986), who reports that co-wives insult and denigrate each other to gain the husband's attention. In a study of the psychological consequences of polygyny among beduin Arabs in the Negev, Al-Krenawi and Graham (2006) report that co-wife conflict is almost universal and results in higher psychological distress, somatisation, phobia, and lower marital and life satisfaction for females, as well as poor family functioning in general. Jankoviak et al. (2005) study a sample of 69 cultures in which polygyny is common, and find no incidence of harmonious co-wife relations. At best there is pragmatic cooperation but more common are severe co-wife conflicts because of a struggle for resources as well as sexual and emotional discomfort with sharing the same husband[7].

3 Theoretical Framework

In this section, we develop a model of fertility and child survival. Formalizing the qualitative anthropological evidence, we derive conditions for the efficiency of observed fertility and child mortality outcomes across households with different observable characteristics.

THE SETTING

Consider a household with $I \geq 1$ wives and a vector of other observable characteristics x, for example the literacy status of the husband and each of his wives. Each wife is indexed by i according to her rank ($i = 1, 2, ...$). Given the prevailing systems of inheritance and household partition at the time of the head of household's death, each wife derives utility from only her own surviving physical offspring. This utility is denoted by $u_i(m_i^b, m_i^g)$, where $m_i^k,\ k = b, g$ denotes the

[7] Polygyny and co-wife conflict in a western context is portrayed by Josephson (2002) in a study of polygyny among early Mormon settlers in Utah.

number of surviving children of sex k, and b and g denote boys and girls, respectively. We assume that u is increasing and concave in both arguments.

A male (female) child survives with probability p_i^b (p_i^g), where survivals of children of a given mother are statistically independent events. Providing p_i^k entails a private cost to the physical mother, which may also depend on her total fertility, n_i. This captures the observation that additional fertility depletes a mother's health (Kotwal, 2002). Moreover, this cost may depend on a mother's personal characteristics relevant to providing child care summarized in the scalar γ_i. To be precise, we let $c^k(p_i^k, n_i)/\gamma_i$ denote the cost of providing survival probability p_i^k to a child of sex k, where c^k is increasing in both arguments. We shall, moreover, assume that c^k is convex in its first argument, which captures the feature of decreasing marginal returns (in terms of survival) to resources devoted to a child. The parameter γ_i captures, first, a mother's efficiency for providing child care and, second, the genetically inherent fitness of wife i's offspring. When we refer to γ_i as reproductive fitness in the sequel, we imply both of these attributes. Notice that, the larger γ_i, the smaller the cost of providing a given level p_i^k conditional on n_i. Further, while the functions $c^b(p_i^b, n_i)$ and $c^g(p_i^g, n_i)$ may differ, the functional form through which γ_i affects the cost of providing p_i^k captures the notion that, *ceteris paribus*, higher reproductive fitness benefits the survival chances of sons and daughters similarly.

We next turn to household decisions regarding fertility. We will consider two alternative assumptions. First, in accordance with Strassmann's (1997) descriptions and other accounts from Sub-saharan Africa cited in Tertilt (2005), it will be assumed that a wife's fertility n_i is controlled by her husband and thus exogenous to her. Second, we will consider the case where a wife is able to control her fertility. In light of the narrative evidence, we consider the former assumption as the more relevant one. To add robustness to the subsequent empirical analysis, however, we will also consider the latter in some detail.

Total resources available to a wife for child rearing, which are denoted by y_i and which may include non-pecuniary parental inputs provided by the father, are assumed to be stipulated by the household head and thus to be exogenous to a wife. According to the qualitative evidence, the majority of rural women conduct their own agricultural as well as non-agricultural businesses and are entitled to the resulting income themselves. The head of household, however, typically also contributes some of his personal income toward child rearing, of which one may think as a transfer to the wife. Given that he can observe the wife's own earnings, he can control her eventual budget by choosing the amount of this transfer accordingly.

We will assume, on the other hand, that a wife has control over how she allocates resources across her children. This is in accordance with the narrative evidence that typically each wife in a polygynous union has her own kitchen and administers her own sub-household, while the household head's involvement in parenting is rather limited. In terms of the analytical framework set out above, wife i chooses p_i^g and p_i^b, the survival chances of her female and male offspring, respectively.

To ease exposition and without loss of generality, we consider a wife with a balanced sex ratio among her children, i.e. she gives birth to $n_i/2$ boys and $n_i/2$ girls. Children of the same sex born to a given wive are assumed to be identical in terms of survival-relevant characteristics. To keep the analysis tractable, we abstract from complexities introduced by the sequential nature of child bearing and mortality, and make the simplifying assumption that all children are born simultaneously.

HUSBAND CONTROLS FERTILITY

In this scenario, at the first stage, the husband chooses a fertility and resource allocation $(n_1, ..., n_I, y_1, ..., y_I)$, which is exogenous to each wife. At the second stage, each wife maximizes her expected utility subject to a resource constraint by choice of p_i^b and p_i^g. In the third stage, child survival is realized. The properties of u and c^k imply that a mother will choose an identical survival probability of p_i^b for all boys and an identical survival probability of p_i^g for all girls. In general, the survival probabilities chosen for boys and girls may, of course, differ.

At the second stage, wife i's maximization problem is

$$\begin{aligned} & \max_{p_i^b, p_i^g} E\left[u_i(M_i^b, M_i^g)\right] \qquad (1) \\ = \; & \max_{p_i^b, p_i^g} \sum_{h=0}^{n_i/2} \sum_{j=0}^{n_i/2} \left\{ u_i(h,j) \begin{pmatrix} n_i/2 \\ h \end{pmatrix} \left(p_i^b\right)^h \left(1-p_i^b\right)^{n_i/2-h} \begin{pmatrix} n_i/2 \\ j \end{pmatrix} \left(p_i^g\right)^j \left(1-p_i^g\right)^{n_i/2-j} \right\} \end{aligned}$$

subject to

$$\frac{n_i}{2}\left(c^b(p_i^b, n_i)/\gamma_i + c^g(p_i^g, n_i)/\gamma_i\right) \leq y_i. \qquad (2)$$

Notice that M_i^b and M_i^g are distributed independently as binomial random variables with parameters p_i^b and $n_i/2$, and p_i^g and $n_i/2$, respectively.

The relevant part of the solution to this constrained maximization are the two demand functions

$$p_i^{k*}(n_i, \gamma_i, y_i), \; k = b, g, \qquad (3)$$

which express the chosen survival probability as a function of fertility, the mother's reproductive fitness, and her resource endowment. Notice that, by virtue of (2), p_i^{k*} is homogenous of degree zero in y_i and γ_i^{-1}, i.e.

$$p_i^{k*}(n_i, \gamma_i/\lambda, \lambda y_i) = p_i^{k*}(n_i, \gamma_i, y_i). \tag{4}$$

WIFE CONTROLS FERTILITY

In this scenario, at the first stage, the husband chooses a resource allocation $(y_1, ..., y_I)$, which is exogenous to each wife. At the second stage, each wife maximizes her expected utility subject to a resource constraint by choice of n_i, p_i^b and p_i^g. In the third stage, child survival is realized. Wife i's maximization problem is the same as in (1), except that she has an additional choice variable, n_i.

The relevant part of the solution are the two demand functions

$$\widetilde{p}_i^{k*}(\gamma_i, y_i),\ k = b, g, \tag{5}$$

where, as before, $\widetilde{p}_i^{k*}$ is homogenous of degree zero in y_i and γ_i^{-1}.

COMPARATIVE STATICS AND TESTABLE IMPLICATIONS

We are interested in comparative static properties of p_i^{b*} and p_i^{g*} when the vector of household traits x changes. In the subsequent discussion, we focus on the case where the household head controls fertility and turn to the alternative scenario toward the end of this section.

Within our framework, a change in x may affect p_i^{k*} through three channels, fertility, reproductive fitness, and the resources available to a wife for child rearing. A literate husband, for example, may choose a smaller number of children than an illiterate one. The former may also devote a larger budget for child care to each wife. Moreover, a literate husband may on average be matched with a woman of different reproductive fitness, which remains unobserved by the researcher. In what follows, we will pursue an approach that remains agnostic about how x affects n_i, γ_i and y_i. As is common in the empirical literature on the collective household model, we maintain the assumption that u's marginal rate of substitution is unaffected by a change in x.

Denoting by x_l the l'th trait relevant to wife i ($l = 1, ..., L$), from (3) we obtain

$$\frac{dp_i^{k*}}{dx_l} = \frac{\partial p_i^{k*}}{\partial y_i}\frac{dy_i}{dx_l} + \frac{\partial p_i^{k*}}{\partial \gamma_i}\frac{d\gamma_i}{dx_l} + \frac{\partial p_i^{k*}}{\partial n_i}\frac{dn_i}{dx_l} \tag{6}$$

Moreover, by virtue of (4), $\frac{dp_i^k}{dy_i}$ and $\frac{dp_i^k}{d\gamma_i}$ are proportional and we may simplify (6) to

$$\frac{dp_i^{k*}}{dx_l} = \delta_{il}\frac{\partial p_i^{k*}}{\partial y_i} + \frac{\partial p_i^{k*}}{\partial n_i}\frac{dn_i}{dx_l} \text{ for } k = b, g \text{ and } l = 1, ..., L, \tag{7}$$

where $\delta_{il} = \left(\frac{dy_i}{dx_l} + \frac{y_i}{\gamma_i}\frac{d\gamma_i}{dx_l}\right)$, which is independent of k. Solving (7) with $k = b$ for δ_{il} and substituting into (7) with $k = g$, we obtain

$$\frac{dp_i^{g*}}{dx_l} = \pi_i\frac{dp_i^{b*}}{dx_l} + \left(\frac{\partial p_i^{b*}}{\partial n_i} - \pi_i\frac{\partial p_i^{g*}}{\partial n_i}\right)\frac{dn_i}{dx_l}, \text{ where } \pi_i = \frac{\frac{\partial p_i^{b*}}{\partial y_i}}{\frac{\partial p_i^{g*}}{\partial y_i}}. \tag{8}$$

Notice that π_i is independent of l. Moreover, $\frac{dp_i^{g*}}{dx_l}, \frac{dp_i^{b*}}{dx_l}, \frac{\partial p_i^{b*}}{\partial n_i}, \frac{\partial p_i^{g*}}{\partial n_i}$ and $\frac{dn_i}{dx_l}$ are effects which can be estimated from data on child survival and fertility. In particular, $\frac{\partial p_i^{g*}}{\partial n_i}$ is a daughter's survival response to an increase in fertility, while $\frac{dn_i}{dx_l}$ is the change in fertility in response to a change in x_l. $\frac{\partial p_i^{b*}}{\partial y_i}$ and $\frac{\partial p_i^{g*}}{\partial y_i}$, on the other hand, remain unobserved by the researcher. As only the ratio between these two terms matters, however, data from households which differ in L traits allow a test of $L-1$ independent restrictions per wife.

Each restriction relates the change in girls' to the change in boys' survival chances in response to a change in trait x_l. To provide intuition, suppose neither x_1 nor x_2, e.g. literacy status of the mother and father, affects the mother's fertility, hence $\frac{dn_i}{dx_1} = \frac{dn_i}{dx_2} = 0$. Then (8) implies that girls' and boys' survival chances change proportionally (with proportionality factor π_i) in response to a change in any of the two traits. This is because a change in either of the traits, according to the model, may only result in a change in reproductive fitness or the budget available for child rearing. Therefore, if a change in the father's literacy status, say, changes the survival probabilities of boys twice as much as for girls, the same holds for a change in the mother's literacy status. When a trait also affects fertility choices, the proportional relationship between girls' and boys' survival chances is "corrected" by the differential impact of the resulting fertility change on mortality.

When each wife makes her own fertility decision, instead of (6) we have

$$\frac{d\widetilde{p}_i^{k*}}{dx_l} = \frac{\partial \widetilde{p}_i^{k*}}{\partial y_i}\frac{dy_i}{dx_l} + \frac{\partial \widetilde{p}_i^{k*}}{\partial \gamma_i}\frac{d\gamma_i}{dx_l}$$

and the set of testable restrictions becomes

$$\frac{d\widetilde{p}_i^{g*}}{dx_l} = \widetilde{\pi}_i \frac{d\widetilde{p}_i^{b*}}{dx_l}, \text{ where } \widetilde{\pi}_i = \frac{\frac{\partial \widetilde{p}_i^{b*}}{\partial y_i}}{\frac{\partial \widetilde{p}_i^{g*}}{\partial y_i}}. \tag{9}$$

Thus, when a wife chooses fertility herself, the survival probabilities for girls and boys change exactly proportionally in response to a change in a household trait. As before, data on households differing according to L traits generate $L-1$ independent restrictions per wife.

4 Econometric Specifications

As in the preceding section, we first focus on the scenario where the household head controls fertility. The set of restrictions (8) can be tested within a joint econometric model of fertility and child mortality. The model consists of two estimation equations, one with fertility, and one with mortality as the dependent variable. The traits used as explanatory variables in both equations are literacy measures for husband and wife as well as a household wealth indicator. The mortality equation allows for different effects of the independent variables according to the child's sex.

Fertility

We set out the fertility equation first. We model the spacing between any two consecutive deliveries of a wife. This duration, W_{jik} say, is modelled within a proportional hazard framework.[8] We let j index households, i the rank of wives within a household, and k the spells between deliveries. To be precise, for all births that have occurred before the interview date, W_{jik} equals the time (in months) between the k'th and $k+1$'th birth. If a woman reports K births at the interview, W_{jiK} is an uncompleted spell and equal to the time (in months) between the woman's last delivery and the interview date.

Let $\lambda_{jik}^F(w)$ denote the birth hazard for child $k+1$ in month w, where w is the time since the k'th delivery. According to the proportional hazard framework, we specify

$$\log \lambda_{jik}^F(w) = \log \lambda_0^F(w, \alpha^F) + \beta^F x_{ji}^F, \tag{10}$$

where $\lambda_0^F(w, \alpha^F)$ is the baseline hazard, which is a function of a parameter vector α^F, β^F is a

[8] A proportional hazard approach to birth spacing has previously been used by Makepeace and Pal (2005).

row vector of coefficients, and x^F_{ji} is a vector of explanatory variables. We restrict attention to characteristics which do not vary over births for a given wife. Therefore x is not indexed by k. Notice that a positive sign of the s'th element of β^F implies that the birth hazard and thus fertility is increasing in the s'th element of x^F_{ji}.

The log-likelihood for a completed spell of duration w_{jik} equals W_{jik}'s log-density, which is

$$\log L^F_{jik} = \log \lambda^F_0(w_{jik}, \alpha^F) + \beta^F x^F_{ji} - e^{\beta^F x^F_{ji}} \Lambda^F_0(w_{jik}, \alpha^F),$$

where $\Lambda^F_0(w_{jik}, \alpha^F)$ denotes the integrated baseline hazard,

$$\Lambda^F_0(w, \alpha^F) = \int_0^w \lambda^F_0(t, \alpha^F) dt.$$

For the uncompleted spell W_{jiK}, the likelihood equals the probability of the $K+1$'th delivery not having occurred by the time of the interview, which gives

$$\log L^F_{jik} = -\Lambda^F_0(w_{jik}, \alpha^F) e^{\beta^F x^F_{it}}.$$

The log-likelihood function for wives of rank i is thus

$$\mathscr{L}^F_i = \sum_{j=1}^{J} \left[\sum_{k=1}^{K_{ji}-1} \left(\log \lambda^F_0(w_{jik}, \alpha^F) + \beta^F x^F_{it} \right) - \sum_{k=1}^{K_{ji}} e^{\beta^F x^F_{it}} \Lambda^F_0(w_{jik}, \alpha^F) \right], \tag{11}$$

where K_{ji} denotes the number of deliveries of wife i of household j which have occurred before the interview.

As in Makepeace and Pal (2005), we specify the baseline hazard as a step function with two nodes. With ten months as minimum spacing when no premature delivery occurs, we choose 16 and 28 months as nodes. To be precise,

$$\lambda^F_0(w, \alpha^F) = \begin{cases} \alpha^F_0, \text{ if } w \leq 16 \\ \alpha^F_0 + \alpha^F_1, \text{ if } 16 < w \leq 28 \\ \alpha^F_0 + \alpha^F_1 + \alpha^F_2, \text{ if } 28 < w. \end{cases}$$

This specification accommodates for the possibility that, conditional on observable characteristics, the birth hazard may depend on time since previous birth.

MORTALITY

As for fertility, we adopt a proportional hazard framework[9]. We define as the dependent variable the survival time of a child up to her/his fifth birthday, T_{jik}. To be precise, for the $k+1$'th child of wife i in household j, T_{jik} is equal to the number of survived months if the child is dead at the time of the interview. In this case, T_{jik} is a completed spell. When the child was born more than five years ago and has survived at least five years, $T_{jik} = 60$ and the observation is treated as an uncompleted spell. When the child is born less than five years before the interview and still alive at the time of the interview, T_{jik} equals the child's age (in months) at the time of the interview, and the observation is treated as an uncompleted spell as well.

Denoting by $\lambda^M_{jik}(s)$ the mortality hazard of the $k+1$'th child of wife i in household j at age s (in months), it is assumed that

$$\log \lambda^M_{jik}(s) = \log \lambda^M_0(s, \alpha^M) + \beta^M x^M_{jik}.$$

As x includes the spacing between the k'th and $k+1$'th birth, first births will not be used in this analysis. Notice that a positive sign of the s'th element of β^M implies that the death hazard, and thus mortality, is increasing in the s'th element of x^M_{jik}.

As for fertility, we specify the mortality baseline hazard as a step function with two nodes, which are chosen at six and twelve months. Also as before, completed spells contribute the probability density to the likelihood function, while uncompleted spells contribute the probability of surviving for 60 months or up to the interview date, whichever is shorter. Define by $\varkappa_{ij}$ the subset of indices k for which a child of wife i in household j is reported as dead at an age younger than 60 months. For wives of rank i this gives

$$\mathscr{L}^M_i = \sum_{j=1}^{J} \left[\sum_{k \in \varkappa_{ij}} \left(\log \lambda^M_0(t_{jik}, \alpha^M) + \beta^M x^M_{jik} \right) - \sum_{k=1}^{K_{ji}-1} e^{\beta^M x^M_{jik}} \Lambda^M_0(t_{jik}, \alpha^M) \right]. \tag{12}$$

Notice that $k = 1$ refers to the second child of a wife and $k = K_{ji} - 1$ to the last child born to her before the interview[10].

[9] For child survival data, this econometric model has been used previously by Arney (2002), Kovsted et al. (2003), and many more.

[10] We have also experimented with several alternative econometric specifications for the fertility and mortality equations, e.g. lognormal birth spacing and mortaliy probit (Bhalotra and van Soest, 2006), and fertility probit and mortality probit (Pitt, 1997). We found, however, that none of these specifications fitted the duration data as satisfactorily as the two equation proportional hazard model.

JOINT ESTIMATION AND IMPLEMENTATION OF TESTABLE RESTRICTIONS

For wives of rank i, the joint fertility-mortality log-likelihood is obtained as

$$\mathscr{L}_i = (L)_i^F + (L)_i^M, \tag{13}$$

which is maximized over the parameter vectors $\alpha^F, \alpha^M, \beta^F, \beta^M$ to obtain unrestricted estimates[11]. The testable implication of the efficiency hypothesis in (8) rests on differential effects of household traits on girls' versus boys' survival chances. We therefore partition the vectors β^M and x^M_{jik} according to the sex of the child for the traits of interest. In a bigynous household, those are mother's literacy, Lit_{ji}, co-wife's literacy, $LitO_{ji}$, husband's literacy, $LitM_{ji}$, and household wealth, $Wealth_{ji}$. To be specific, we write

$$\begin{aligned}\beta^M x^M_{jik} \quad = \quad & Boy_{jik} \left\{ \begin{array}{c} \beta_0^{M,b} + \beta_1^{M,b} Lit_{ji} + \beta_2^{M,b} LitO_{ji} + \beta_3^{M,b} LitM_{ji} \\ + \beta_4^{M,b} Wealth_{ji} + \beta_5^{M,b} Spacing_{jik} \end{array} \right\} \\ & + (1 - Boy_{jik}) \left\{ \begin{array}{c} \beta_0^{M,g} + \beta_1^{M,g} Lit_{ji} + \beta_2^{M,g} LitO_{ji} + \beta_3^{M,g} LitM_{ji} \\ + \beta_4^{M,g} Wealth_{ji} + \beta_5^{M,g} Spacing_{jik} \end{array} \right\} \\ & + \beta^{M,c} x^{M,c}_{ji}. \end{aligned} \tag{14}$$

Boy_{jik} is a dummy variable equal to one when child $k+1$ is male and zero otherwise. $Spacing_{ji}$ equals the spell since the last birth and has been denoted by w_{jik} in the statistical derivations. $x^{M,c}_{ji}$ is a vector of control variables, including age and ethnicity of the mother, and $\beta^{M,c}$ is the subset of coefficients in β^M parametrizing those controls.

Differentiation by sex of child is, of course, not meaningful for the fertility equation. Otherwise we parametrize accordingly,

$$\beta^F x^F_{ji} = \beta_0^F + \beta_1^F Lit_{ji} + \beta_2^F LitO_{ji} + \beta_3^F LitM_{ji} + \beta_4^F Wealth_{ji} + \beta_5^F Spacing_{jik} + \beta^{F,c} x^{F,c}_{ji}. \tag{15}$$

To test the efficiency hypothesis for wives of rank i according to equation 8, notice the following correspondences between terms in (8) and parameters of the econometric model as spelled out in

[11] We do not attempt to address the potential fertility selection problem, which may, in principle, bias the coefficients of the mortality equation. See Pitt (1997) for a discussion. He finds, however, that the empirical relationship between measures of literacy and child mortality remains virtually unchanged in data from each of 14 African countries when fertility selection is not accounted for.

(14) and (15),

$$\begin{array}{cc}
\frac{dp_i^{g*}}{dx_l}: & \beta_1^{M,g},...,\beta_4^{M,g} \\
\frac{dp_i^{b*}}{dx_l}: & \beta_1^{M,b},...,\beta_4^{M,b} \\
\frac{\partial p_i^{g*}}{\partial n_i}: & \beta_5^{M,g} \\
\frac{\partial p_i^{b*}}{\partial n_i}: & \beta_5^{M,b} \\
\frac{dn_i}{dx_l}: & \beta_1^F,...,\beta_4^F.
\end{array}$$

This leads us to the following econometric version of the efficiency test (8),

$$\beta_l^{M,g} = \pi_i \beta_l^{M,b} + (\beta_5^{M,b} - \pi_i \beta_5^{M,g})\beta_l^F,\ l = 1,...,4. \tag{16}$$

The test is implemented by maximizing (13), first unrestricted and, second, subject to (16), where π_i is estimated. By standard arguments, twice the difference between the unrestricted and restricted log-likelihood is asymptotically distributed as a Chi-Square statistic with degrees of freedom equal to the number of traits used in (16) minus one.

When each wife chooses fertility herself, spacing becomes a choice variable and the demand for child survival is not causally dependent on fertility. In this case, the appropriate way to implement the efficiency test (9) is to estimate the mortality equation without *Spacing* as an explanatory variable and impose the restrictions

$$\beta_l^{M,g} = \pi_i \beta_l^{M,b},\ l = 1,...,4. \tag{17}$$

This test is simpler to implement as it involves estimation of only one equation and hence no cross-equation restrictions are involved.

IDENTIFICATION

As is clear from section 3, the set of restrictions (8) is valid in general only under the model's particular assumptions. The strength of our identification strategy is that it differences out unobserved changes in reproductive fitness or differential access to resources which are potentially correlated with observable traits. Therefore our test is robust to selection on unobserved reproductive fitness in the marriage market. To illustrate, Strassmann (1997) concludes that polygyny is a risk factor to child survival among Mali's Dogon. Her inference is based on a regression of child survival on observable household and parental characteristics, and a polygyny dummy. A negative estimate of this latter regressor motivates her conclusions. Such an approach can be flawed by several complications. First, the decision to marry a second wife may be driven by the realization of the first wife's reproductive fitness if this characteristic is unobserved at the time of marriage. Similarly, a first wife of high reproductive fitness may be in a better position to hinder her husband to take an additional wife. Second, if reproductive fitness is observed before marriage within the population but unobserved by the researcher, women of lower reproductive fitness may be more likely to be matched into polygynous unions when monogamy is deemed more desirable by women, which Strassmann (1997) indeed reports.

Our framework remains agnostic about differences in child mortality outcomes across different categories of wives. Within the model just developed such differences can in fact be explained by unobserved differences in reproductive fitness of women who get matched into different categories of marital unions. Instead, the key feature of our identification strategy is the focus on the difference between sons' and daughters' outcomes and exploiting the fact that the sex composition of children born to a mother is exogenous to her, at least in the environment studied, where sex-selective abortion is uncommon.

The remaining concern about robustness of our test is that (8) and (9) rest on the assumption of the mother's preferences (or more precisely the marginal rate of substitution between surviving boys and girls) remaining unchanged in response to changes in a trait. While constant preferences are a common assumption in much of the extant empirical work on unitary and collective household models (Browning and Chiappori, 1998; Browning et al., 1994; Maluccio and Quisumbing, 2003), it is, in our view, a somewhat disconcerting assumption to make.[12]. In the discussion that follows we will address how each of the four traits that we have suggested may work as a preference shifter. In particular, we will distinguish between direct and indirect potential effects of each of these traits on preferences.

A most immediate direct effect is likely to arise from changes in the wife's own literacy status. In this connection, it is conceivable that a literate wife has a different relative appreciation of daughters than an illiterate one, especially when education involves family planning.

In contrast, $LitM$, $LitO$, and $Wealth$, whose realization is largely accounted for by the household head given the prevailing bride-price system, likely affect the mother's preferences less directly. Two remaining conceivable channels are, however, selection and adjustment. To give an example, selection is at work when a literate husband matches with a type of wife that appreciates daughters relatively more than the type of wife an illiterate husband is matched with. Second, wives' preferences, even if initially of the same type, may adjust to the respective husband's pref-

[12] While (8) and (9) continue to hold as long as all traits used in the test affect a wife's preferences in the same way, this is an unlikely fluke.

erences, which in turn may be a function of traits such as $LitM$ and $Wealth$.

The way we partially address this problem in our empirical analysis is to consider monogamous households as a comparison group. While it is not a given that allocations are efficient there, monogamous households do not face the issue of potentially resource consuming co-wife rivalry. Moreover, existing studies from largely monogamous societies have so far produced no evidence against the collective household model (e.g. Maluccio and Quisumbing, 2003).

5 Empirical Results

To facilitate comparisons between monogamous and polygynous regimes, we conduct all estimations and tests separately for (i) children of wives in monogamous households, (ii) children of senior wives in bigynous households, and (iii) children of junior wives in bigynous households. Among polygynous unions, we restrict attention to bigynous households because, first, this is by far the most common form of polygyny in the sample and, second, it keeps the polygynous sample as homogenous as possible. Moreover, in the estimations for wives in bigynous households we use only births that occurred while the household head has had two wives, i.e. children born to the senior wife while she was the household head's only wife, are discarded. We first discuss estimation results for the fertility and mortality equations and then turn to the efficiency test.

Estimation

The results of an unrestricted estimation of (11) are set out in Table 2. As far as the traits of interest are concerned, some interesting differences between monogamous and bigynous households arise. In partiuclar, female literacy has no significant on monogamous and senior bigynous wives' fertility while it increases junior bigynous wives' fertility. Male literacy does not significantly affect senior bigynous wives' fertility while it significantly decreases the birth hazard of monogamous as well as junior bigynous wives. Household wealth, on the other hand, has an almost identical negative effect on fertility in all categories. Across all three categories, fertility has significantly decreased between the 1996 and 2001 DHS waves, most dramatically so, however, for bigynous wives.

The results of an unrestricted estimation of (12), the child mortality model, are set out in Table 3. It should be noted upfront that each coefficient estimate quantifies a reduced form, not a causal, effect of a wife or household level trait on child mortality. To be precise consider, for example, wife's literacy. Within our framework, a change in this trait may be associated not only with higher reproductive fitness due to additional learned skills, but also with higher reproductive fitness due to characteristics genetically inherent to the wife, or with a different type of husband due to matching processes in the marriage market. A causal effect, on the other hand, would have to be net of these latter two factors. The efficiency tests (16) and (17), however, are based on restrictions regarding observable reduced-form effects.

As for fertility, some remarkable differences between senior and junior wives in bigynous households arise. Female literacy results in lower child mortality for junior wives for boys but not for girls. Literacy affects the outcome of neither boys nor girls for senior wives, in contrast. What this suggests is that literate junior wives allocate more resources to their sons, at the expense of their daughters. When a junior wife is matched with a literate instead of an illiterate senior wife, in contrast, her sons fare worse. The difference between the effects on the junior wife's sons and daughters is, however, not statistically significant. Household wealth as well as husband's literacy significantly reduces the mortality of senior wives' sons and junior wives' daughters.

For monogamous households, in contrast, the results are overall much more balanced across boys and girls. Each of the three traits available for monogamous households produces the same direction of effect for sons and daughters and only in the case of wealth is there a significant difference, in favor of girls, between sons and daughters. These findings suggest that, at least within monogamous households, changes in different traits do not trigger substantially different changes in preferences for boys' versus girls' survival. We take this as evidence against the concerns raised in the previous subsection about preference-shifting induced by changes in traits.

Table 4 gives estimates for the mortality equation under the alternative assumption that each wife controls her fertility. Recall that, in this case, spacing is excluded as an explanatory variable. Even though spacing is an important predictor of mortality in the previous specification, none of the coefficients reported in Table 3 changes in a remarkable fashion.

According to the child mortality results, the findings for bigynous households are roughly consistent with a story of struggle between co-wives for the survival of physical sons: a literate junior wife succeeds in channelling more resources to her sons' survival while improvements in household-level status variables which are not wife-specific, husband's literacy and wealth, enable the senior wife to provide her sons an advantage, which the sons of her junior co-wife fail to enjoy. This may be due to the higher status the senior wife enjoys inside the household. Moreover, a literate senior wife reduces the survival chances of all of her co-wife's children while there is no benefit for any of her own offspring, which is consistent with resources being expended unproductively by a literate senior wife to reduce survival chances of rivaling co-wife offspring.

Our theoretical framework is, however, agnostic about such cross-wife dynamics and can, for example, immediately accomodate the finding that more household wealth increases junior wives' survival of daughters by much more than that of her sons. The theoretical analysis implies, however, that then girls' survival is much more elastic in the change in any trait than boys' survival. Similarly, when the effect of senior wife's literacy on survival of children of the junior wife is concerned, within our framework, it may be efficient for a household to divert more resources to a literate senior wife as she may be of higher reproductive fitness. The larger coefficient for boys then, however, suggests that boys' survival is more elastic than girls' (Table 3, column 3, rows 2 and 11), which is at odds with the effects of wealth (column 3, rows 3 and 12) and head of household's literacy (column 3, rows 4 and 13) on mortality of junior wives' children. These findings set the stage for the efficiency tests, which are presented in the next subsection.

Efficiency Tests

For the case where fertility decisions are taken by the husband, which is our preferred specification, we estimate the fertility and mortality models jointly with alternative sets of restrictions (16) imposed. The results are set out in Table 5. For monogamous households, none of the four permutations of traits gives a rejection. Notice that, for monogamous households, there are roughly five times as many observations as for senior or junior wives in bigynous households, i.e. the test has more than twice the power for monogamous households. This finding suggests, therefore, that our model of fertility and efficient child investment is not inappropriate in describing decision making processes in monogamous households. In particular, the failure to reject the null hypothesis lends indirect support to the maintained assumption that changes in traits do not change a wife's preferences in a significant fashion.

The same holds true for senior wives of bigynous households. For junior wives, in contrast, the null hypothesis is rejected with a p-value of four per cent for one of the eight specifications, and with p-values smaller than ten per cent for an additional two specifications. These findings suggest that, while senior wives achieve efficient child survival outcomes, junior wives do not. This is in accordance with anthropological evidence, according to which it is particularly the offspring of the junior wife, the adult with typically the weakest bargaining position in the household, who suffers most severely from co-wife rivalry and competition for own surviving offspring.

Test results for the case where the mother makes fertility decisions and the restrictions are given by (17), are set out in Table 6. The findings of Table 5 are qualitatively confirmed throughout. For monogamous wives and senior bigynous wives, none of the test statistics attains a p-value of ten per cent or smaller while this is the case for junior bigynous wives for six of the eight permutations of traits. The finding of inefficient allocations among junior wives' children is thus robust to the nature of the underlying decision making process regarding fertility.

Do these findings prove that polygyny is responsible for inefficient allocations in a causal sense? Strictly speaking, the answer is negative because, as pointed out in the preceding exposition, matching of women into polygynous unions is self-selected and could thus, in principle, also drive our results.

6 Conclusion

Previous anthropological research on polygyny in Sub-saharan Africa has found that children in polygynous unions are generally at a greater mortality risk, despite of higher fertility of women in monogamous unions. The empirical results of this study complement and qualify these existing findings. First, the variability of fertility and mortality outcomes is greater for wives who marry into the household later. Second, the outcome for boys, in particular those of junior wives in bigynous households, are more variable than for girls. Most importantly, however, we reject a set of restrictions implied by efficient resource allocations among children for junior, but not for senior wives in polygynous households. This confirms and extends two elements of the existing narrative evidence on polygynous households. First, our findings are in accordance Strassmann (1997) and others, who have identified co-wife rivalry as a risk factor for child health in polygynous unions. Second, several accounts relate that junior wives are the adults with the weakest position inside a polygynous household. While previous anthropological research has found that senior wives enjoy non-material privileges, our results suggests that junior wives are also at a significant material disadvantage. Moreover, our finding that the hypothesis of efficient allocations is not rejected for monogamous households provides additional evidence that polygynous households are different from monogamous ones when child survival is concerned.

Our results challenge the collective view of the household, according to which intra-household allocations are efficient. In all empirical applications from developed countries, where the classical unitary model is rejected, the collective model is not. Our findings, in contrast, complement work by Udry (1996) and Duflo and Udry (2005), who use production and expenditure data from households in Burkina Faso and Cote d'Ivoire to reject the collective model. Their approach identifies inefficiencies by exploiting separate, gender-specific entitlements to land within a farm household. This paper, instead focuses on child mortality and exploits differences in survival chances of boys and girls of the same wife in response to changes in household characteristics, to arrive at similar conclusions.

In a recent macroeconomic analysis, Tertilt (2005) has shown that in a male dominated society polygyny can be responsible for excessive fertility, crowd out investment into physical capital, and thus slow down economic growth. Our analysis provides empirical evidence for a welfare loss on the micro level generated by polygyny. Taken together, these papers challenge previous research which has deemed polygyny potentially efficiency-enhancing in environments with a surplus of brides and missing labor markets (Becker, 1994; Jacoby, 1995).

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Table 1: Descriptive Statistics

	All households				Monogamous households	Bygynous households		
						Senior Wives		Junior Wives
	Mean	Std	Min	Max	Mean	Mean		Mean
Wives	N=13,866				N=7,359	N=1,806		N=1,806
Number of Children at interview	4.627	3.155	0	17	4.143	6.229		3.965
Age at interview	30.540	8.855	15	49	28.712	35.372		28.591
Age at time of marriage	16.008	2.790	8	37	16.056	15.717		16.119
Age at delivery	24.812	7	15	49	24.221	28.019		24.566
Year of marriage	1984	9.312	1958	2001	1986	1979		1986
Living in polygynous union	0.469	0.499	0	1	0.000	1.000		1.000
Literacy	0.090	0.287	0	1	0.092	0.072		0.096
1996 survey	0.402	0.490	0	1	0.397	0.435		0.436
Husbands	N= 2,706				N=1,899		N=720	
Number of Children at interview	5.939	4.540	0	28	4.396		9.156	
Age at interview	39.660	9.697	18	59	37.815		43.692	
Age at marriage	24.756	5.094	10	56	24.902		24.525	
Year of marriage	1984	9.913	1955	2001	1986		1980	
Living in polygynous union	0.355	0.479	0.000	1.000				
Wealth (Index) **	-0.363	0.412	-1.124	4.419	-0.381		-0.318	
Literacy	0.173	0.378	0	1	0.190		0.133	
1996 survey	0.402	0.490	0	1	0.404		0.396	
Children	N=6,4312				N=27,952	N=6,198*		N=5,711
Female	0.492	0.500	0	1	0.491	0.499		0.493
Prior Spacing (all births except first)	30.357	16.017	9	268	30.460	31.305		31.250
Twin	0.030	0.171	0	1	0.030	0.029		0.034
Age at interview (survivors)	109.965	83.993	0	444	97.838	85.573		98.162
Died before interview at age 5 or younger	0.267	0.443	0	1	0.253	0.264		0.253
Male	0.279	0.448	0	1	0.261	0.263		0.269
Female	0.256	0.436	0	1	0.244	0.265		0.236
1996 survey	0.412	0.492	0	1	0.461	0.512		0.542

* only children born during bygynous regime

** computed according to Filmer and Pritchett (2001)

Table 2: Proportional Hazard Analysis of Fertility

	(1)	(2)	(3)
	Monogamous	Bygynous Households	
		Senior wives	Junior wives
Traits			
Literacy Wife	-0.022	0.029	0.085
	[0.0241]	[0.0587]	[0.051]*
Literacy Cowife		0.043	-0.041
		[0.0536]	[0.056]
Wealth	-0.071	-0.065	-0.076
	[0.0174]***	[0.0366]*	[0.036]**
Literacy Husband	-0.108	-0.062	-0.125
	[0.0216]***	[0.0534]	[0.058]**
Controls			
Age of Wife at Birth	0.064	0.132	0.005
	[0.0069]***	[0.0173]***	[0.017]
Age of Wife at Birth Squared	-0.025	-0.036	-0.015
	[0.0012]***	[0.0028]***	[0.003]***
1996 Survey	0.412	0.457	0.497
	[0.0147]***	[0.0313]***	[0.032]***
Baseline Hazard			
Constant	0.005	0.002	0.009
	[0.0005]***	[0.0006]***	[0.002]***
Node at 16 Months	0.034	0.014	0.067
	[0.0032]***	[0.0035]***	[0.016]***
Node at 28 Months	0.011	0.004	0.026
	[0.0012]***	[0.0012]***	[0.007]***
Observations	27952	6198	5711

Notes: Robust standard errors in brackets; *, **, ***: significant at 90, 95, and 99% significance level, respectively; three ethnicity dummies which are included in the estimation are not reported. For bygynous households, only births during the bygynous regime are used.

Table 3: Proportional Hazard Analysis of Mortality, Husband Makes Fertility Decisions

	(1)	(2)	(3)
	Monogamous	Bygynous	
		Senior wives	Junior wives
Boys			
Literacy wife	-0.065	0.099	-0.334
	[0.0743]	[0.1603]	[0.164]**
Literacy cowife		-0.069	0.247
		[0.1292]	[0.157]
Wealth	-0.042	-0.166	0.014
	[0.0537]	[0.1100]	[0.100]
Literacy husband	-0.248	-0.402	-0.119
	[0.0716]***	[0.1551]***	[0.172]
Age of Wife at Birth	-0.037	-0.077	0.005
	[0.0719]	[0.0365]**	[0.045]
Age of Wife at Birth Squared	0.007	0.014	0.000
	[0.0129]	[0.0061]**	[0.008]
1996 Survey	0.126	-0.108	0.274
	[0.0484]***	[0.0877]	[0.091]***
Twin	1.186	1.675	1.802
	[0.0749]***	[0.1125]***	[0.126]***
Preceding Birth Interval	-0.029	-0.029	-0.021
	[0.0020]***	[0.0029]***	[0.003]***
Girls			
Literacy wife	-0.059	-0.215	0.018
	[0.0766]	[0.1806]	[0.151]
Literacy cowife		0.034	0.114
		[0.1379]	[0.175]
Wealth	-0.183	-0.154	-0.311
	[0.0557]***	[0.1055]	[0.117]***
Literacy husband	-0.124	0.017	-0.114
	[0.0713]*	[0.1351]	[0.175]
Age of Wife at Birth	-0.048	-0.098	-0.063
	[0.0234]**	[0.0423]**	[0.051]
Age of Wife at Birth Squared	0.009	0.018	0.011
	[0.0043]**	[0.0073]**	[0.009]
1996 Survey	0.208	0.085	0.318
	[0.0444]***	[0.0872]	[0.101]***
Twin	1.358	1.191	1.109
	[0.0783]***	[0.1266]***	[0.151]***
Preceding Birth Interval	-0.030	-0.025	-0.025
	[0.0020]***	[0.0029]***	[0.003]***
Constant	-0.006	0.195	0.857
	[1.0049]	[0.7905]	[0.890]
Baseline Hazard			
Constant	0.074	0.126	0.034
	[0.0712]	[0.0674]*	[0.020]*
Node at 6 Months	-0.046	-0.089	-0.023
	[0.0443]	[0.0480]*	[0.014]*
Node at 12 Months	-0.024	-0.028	-0.009
	[0.0225]	[0.0153]*	[0.005]*
Observations	22177	4959	4467

Notes: Robust standard errors in brackets; *, **, ***: significant at 90, 95, and 99% significance level, respectively; three ethnicity dummies which are included in the estimation are not reported. For bygynous households, only births during the bygynous regime are used.

Table 4: Proportional Hazard Analysis of Mortality, Wife Makes Fertility Decisions

	(1)	(2)	(3)
	Monogamous	Bygynous	
		Senior wives	Junior wives
<u>Boys</u>			
Literacy wife	-0.110	0.091	-0.229
	[0.0610]*	[0.1531]	[0.1372]*
Literacy Cowife		-0.069	0.305
	-0.109	[0.1249]	[0.1309]**
Wealth	[0.0440]**	-0.249	-0.069
		[0.1051]**	[0.0886]
Literacy Husband	-0.150	-0.428	-0.025
	[0.0550]***	[0.1487]***	[0.1378]
Age of Wife at Birth	-0.069	-0.095	-0.066
	[0.0184]***	[0.0345]***	[0.0390]*
Age of Wife at Birth Squared	0.010	0.016	0.010
	[0.0035]***	[0.0059]***	[0.0073]
1996 Survey	0.299	0.003	0.468
	[0.0332]***	[0.0811]	[0.0765]***
Twin	1.119	1.588	1.795
	[0.0572]***	[0.1107]***	[0.1139]***
<u>Girls</u>			
Literacy wife	-0.106	-0.102	0.164
	[0.0646]	[0.1631]	[0.1312]
Literacy Cowife		0.004	0.061
		[0.1304]	[0.1583]
Wealth	-0.196	-0.124	-0.278
	[0.0464]***	[0.0976]	[0.1038]***
Literacy Husband	-0.140	0.017	-0.407
	[0.0582]**	[0.1276]	[0.1631]**
Age of Wife at Birth	-0.094	-0.079	-0.057
	[0.0193]***	[0.0388]**	[0.0438]
Age of Wife at Birth Squared	0.015	0.013	0.008
	[0.0037]***	[0.0068]*	[0.0082]
1996 Survey	0.324	0.139	0.396
	[0.0351]***	[0.0802]*	[0.0865]***
Twin	1.322	1.320	0.960
	[0.0590]***	[0.1235]***	[0.1520]***
Constant	0.138	-0.222	-0.232
	[0.3318]	[0.7148]	[0.7395]
<u>Baseline Hazard</u>			
Constant	0.061	0.078	0.051
	[0.0138]***	[0.0380]**	[0.0249]**
Node at 6 Months	-0.039	-0.056	-0.035
	[0.0090]***	[0.0271]**	[0.0172]**
Node at 12 Months	-0.017	-0.017	-0.012
	[0.0040]***	[0.0084]**	[0.0060]**
Observations	27952	6198	5711

Notes: Robust standard errors in brackets; *, **, ***: significant at 90, 95, and 99% significance level, respectively; three ethnicity dummies which are included in the estimation are not reported. For bygynous households, only births during the bygynous regime are used.

Table 5: Efficiency Tests, Husband Makes Fertility Decisions

		Monogamous	Byginous	
Traits (Degres of Freedom)			Senior Wives	Junior Wives
LitM, Lit (1)	Chi-Sq.	0.08	1.34	0.4
	P	0.7735	0.2474	0.5294
Lit, Wealth (1)	Chi-Sq.	0.4	1.53	4.26
	P	0.527	0.216	0.0391
LitM, Wealth (1)	Chi-Sq.	4.16	0.46	1.512
	P	0.1248	0.498	0.219
LitM, LitO (1)	Chi-Sq.		0.05	0.08
	P		0.8301	0.7825
Lit, Wealth, LitM (2)	Chi-Sq.	4.2	3.01	4.79
	P	0.2407	0.2222	0.0912
Lit, LitO, Wealth (2)	Chi-Sq.		1.68	5.55
	P		0.4314	0.0625
LitM, Lit, LitO (2)	Chi-Sq.		1.34	0.59
	P		0.5116	0.7454
LitM, Lit, LitO, Wealth (3)	Chi-Sq.		3.01	5.66
	P		0.3897	0.1293

Table 6: Efficiency Tests, Wife Makes Fertility Decisions

Traits (Degres of Freedom)		Monogamous	Byginous Senior Wives	Junior Wives
LitM, Lit (1)	Chi-Sq.	1.540	0.570	2.960
	P	0.214	0.449	0.086
Lit, Wealth (1)	Chi-Sq.	1.910	0.860	3.870
	P	0.167	0.355	0.049
LitM, Wealth (1)	Chi-Sq.	2.190	1.210	0.460
	P	0.139	0.271	0.500
LitM, LitO (1)	Chi-Sq.		0.240	4.440
	P		0.627	0.035
Lit, Wealth, LitM (2)	Chi-Sq.	2.380	1.610	3.870
	P	0.304	0.446	0.144
Lit, LitO, Wealth (2)	Chi-Sq.		0.870	6.750
	P		0.647	0.034
LitM, Lit, LitO (2)	Chi-Sq.		0.580	6.410
	P		0.749	0.041
LitM, Lit, LitO, Wealth (3)	Chi-Sq.		1.620	7.260
	P		0.655	0.064